Across Time and Space: Variations in Hospital Use During Canadian Health Reform

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Objectives. To investigate change in hospital utilization in a population and to discuss analytical strategies using large administrative databases, focusing on variations in rates of different types of hospital utilization by income quintile neighborhoods.

Data Sources. Hospital discharge abstracts from Manitoba Health, used to study the changes in utilization rates over eight fiscal years (1989–1996).

Study Design. We test the hypotheses that health reform has changed utilization rates, that utilization rates differ significantly across income quintiles (defined by the relative affluence of neighborhood of residence), and that these variations have been maintained over time. Our approach uses generalized estimating equations to produce robust and consistent results for studying rates of recurrent and nonrecurrent events longitudinally.

Data Extraction Methods. Rates of individuals hospitalized, hospital discharges, days of hospitalization, and hospitalization for different types of medical conditions and surgical procedures are generated for the period April 1, 1989 through March 31, 1997 for residents of Winnipeg, Manitoba. Data are grouped according to the individual's age, gender, and neighborhood of residence on April 1 of each of the eight fiscal years for the rate calculations. Neighborhood of residence and the 1991 Canadian Census public use database are used to assign individuals to income quintiles.

Principal Findings. The substitution of outpatient surgery for inhospital surgery accounted for much of the change in hospital utilization over the 1989–1996 period. Health care reform did not have a significant effect on the utilization gradient already observed across socioeconomic groups. Health reform markedly accelerated declines in in-hospital utilization.

Conclusions. Grouping the data with key characteristics intact facilitates the statistical analysis of utilization measures previously difficult to study. Such analyses of variations across time and space based on parametric models allows adjustment for continuous covariates and is more efficient than the traditional nonparametric approach using standardized rates.

Key Words. Health services, longitudinal data; correlation; health reform; socioeconomic status

Substantial variation in the provision of medical and surgical services may reflect system capacity (bed and physician supply), patient health status, a lack of consensus within the provider community, and the discretionary nature of some procedures. In Canada, reductions in acute care beds have been one of the main methods of trying to implement health reform (Tully and Saint-Pierre 1997). Such cost containment efforts suggest the need to study differences in hospital utilization rates before and after a change in the supply of hospital beds. Despite recognition of the importance of various factors in influencing health care utilization, research in both the United States and Canada has demonstrated that individuals of low socioeconomic status "have poor health status and spend many more days in the hospital" (Roos and Mustard 1997; National Center for Health Statistics 1998; Statistics Canada 1994). How has the mandated reduction in bed supply changed such relationships? The longitudinal research reported here follows up on earlier cross-sectional work that investigated how a universally insured health care system delivers care according to socioeconomic characteristics strongly related to health status (Roos and Mustard 1997).

Given the changes under way in the delivery of health care in the Organization for Economic Cooperation and Development (OECD) countries, the development of efficient ways to analyze large data sets across time and space is critical for both epidemiologic and health services research. This article presents a statistical modeling strategy for analyzing large longitudinal data sets, taking into account variation across areas and/or income groups. Our strategy builds on the generalized estimating equations (GEE) method based on the marginal distribution of the event of interest (Liang and Zeger

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1986). The estimating equations give consistent estimates of the regression parameters and their variance under mild assumptions about time dependence. (See the appendix for a brief review of the GEE method.)

In Manitoba, Canada, the mandated reduction in bed supply has emphasized closing beds in Winnipeg, the capital city. Winnipeg has a population of over 600,000, and it is home to the provincial medical school and seven major hospitals. By the end of fiscal 1996, the Winnipeg hospital acute bed count was down to 2,384 from 3,042 at the beginning of 1989; this represented a drop of 22 percent (although 99 nonacute beds were added in this period; see Table 1). Beds were closed across all services in a system that initially appeared to offer little slack. Prior to the cuts, all institutions were operating at occupancy rates greater than 75 percent. Because beds were just beginning to be closed in 1992, we designate fiscal years 1989–1991 as the "pre-reform" and 1992–1996 as the "post-reform" period. We look at changes in hospital care over time for individuals living in neighborhoods of differing socioeconomic status in Winnipeg: whether the variation in rates across income quintiles was significant and maintained over the years. We investigate changes in such rates as individuals hospitalized, overall discharges, and days of hospitalization for short-stay care. In addition to overall utilization, we study low-variation conditions, high-variation medical conditions and surgical procedures, outpatient surgery, and obstetrical hospitalizations. Conditions for which hospitalization might not be necessary are also examined. By comparing fiscal 1992–1996 rates with the utilization rates for the three previous years, both for each quintile and for all quintiles combined, the change in the gradient in utilization between the least affluent and most affluent income quintile neighborhoods can be analyzed.

METHODS

Data Sources

In Manitoba the cost of all medical and hospital care, with minor exceptions, is covered by a government health care plan with no limitations on use except for visits to chiropractors and optometrists. Hospital discharge abstracts from Manitoba Health are used to study the changes in utilization rates over time. Previous work has shown the reporting of hospital use and surgical procedures to be accurate in Manitoba (Roos et al. 1982; Roos, Sharp, and Wajda 1989). We identified all patients admitted as hospital inpatients, excluding newborns (to avoid double counting) and individuals residing out of the province or in a

Table 1:   History of Acute Bed Closures at Winnipeg Hospitals Acute Bed Change 1989–1996

					Fiscal Year				
	1989	1990	1991	1992	1993	1994	1995	1996	Ratio: 96/89
Total Acute Beds at Start of Fiscal Year	3042	3042	3042	3013	2707	2498	2460	2384	0.78
Net Changes by Hospital Type:									
Teaching	0	0	0	−29	−279	−54	−25	−17	
Other	0	0	0	0	−27	−155	−13	−59	
Total	0	0	0	−29	−306	−209	−38	−76	
Change (% decrease)	0	0	0	−1.0	−10.2*	−7.7*	−1.5	−3.1	

Note: These values exaggerate the loss of available beds. Although 306 acute beds closed in 1993, 75 non-acute beds were added to the system; in 1994 despite the closure of another 209 acute care beds, 24 non-acute beds were added. Except for 1993 and 1994, the changes in the number of non-acute beds were very small.

mental institution. Changes from earlier work included excluding individuals under care of the Public Trustees, assigning Indians with treaty status to their place of residence rather than to their band (located outside of Winnipeg), and counting abstracts filed after the official end of the reporting period but submitted in the next year's data. We then constructed a record of all hospital stays by eligible individuals over eight fiscal years, from April 1989 through March 1997. An individual's history began on his or her date of enrollment in the Manitoba Health Insurance Plan and was terminated on death or out-migration.

Following Roos and Mustard (1997), residents of Winnipeg were divided into five equal-size groups based on average neighborhood household income data derived from the 1991 Canadian Census public use database. Data describing characteristics of neighborhood residents, including the percentage of female-headed households, education levels, and unemployment rates, were also taken from the 1991 census. Census data were aggregated at the geographic unit of the enumeration area; the average population of these areas is approximately 700.

Based on 1991 mean household income, the enumeration areas within Winnipeg were ranked from poorest to wealthiest and then grouped into five population quintiles, with each quintile containing 20 percent of the city's population, or approximately 130,000 residents. Postal code at time of admission to hospital was used to designate an individual as a Winnipeg resident. Then, each Winnipeg resident was linked to an enumeration area by residential postal code; for each resident an income quintile rank was assigned, with Q1 designated as the poorest. Small improvements in the definition of income quintile are primarily responsible for minor differences between this study and previously reported results for fiscal 1992 (Roos and Mustard 1997). The updated 1991 quintile definitions were used throughout the study period.

In reporting our results, hospital, surgical, and visit rates were predicted as standardized by age and sex. Denominators were based on counts of individuals in each income quintile in December of each year, with numerators based on event counts (i.e., hospitalizations) for each individual identified as a member of a quintile. The 1991 Winnipeg population was used as the standard population.

Hospital Use

We examine the records of Winnipeg residents hospitalized for all short-stay hospital discharges (1 to 59 days) over the eight years, regardless of where the

hospitalization took place. In fiscal 1996, the proportion of inpatient hospital discharges of Winnipeg residents from Winnipeg hospitals was 95 percent.

Following Wennberg (1986) and Roos and Mustard (1997), more detailed analyses were also carried out for the following categories of medical care:

1. *High-variation surgical conditions* include pediatric and adult inpatient surgical cases with highly variable rates.
2. *Outpatient surgery* includes patients admitted for a valid surgical DRG (diagnostic related group), classified as having either same day or outpatient surgery, and having a length of stay coded as zero days (day of admission noted as the same as day of discharge in Manitoba Health's system).
3. *High-variation medical conditions* apply to conditions, like pneumonia, gastroenteritis, and chronic obstructive lung disease, that have highly variable admission rates. These conditions represent more than 80 percent of medical admissions to the hospital (Wennberg 1986). We analyzed both pediatric and adult discharges.
4. *Low-variation conditions* are medical and inpatient surgical conditions that demonstrate relatively stable rates across populations. In this article, the term applies to acute myocardial infarction, hip fracture, and colon cancer surgery, for which there is little clinical ambiguity about the need for hospitalization. Other analyses have confirmed the appropriateness of this specification (Fisher et al. 1994).
5. *Obstetric hospitalizations* are relevant for women of childbearing age (15–45 years old) delivering newborns in Winnipeg.

Three indicators forwarded as reflecting the need for health care are included: ambulatory care–sensitive conditions (Billings, Zeitel, Lukomnik, et al. 1993); conditions amenable to medical treatment (Charlton et al. 1983; Poikolainen and Eskola 1986; Desmeules and Semenciw 1991); and avoidable hospitalizations (Weissman, Gatsonis, and Epstein 1992). All are based on (1) hospital discharge rates for specific types of conditions for which medical treatment is believed to be effective in preventing the condition; (2) finding and treating the condition in an early phase to avoid major consequences; or (3) treating the condition in a late phase, thereby avoiding death or disability.

Statistical Approach

In reporting the standardized rates of utilization by quintile groups for various conditions, we used the traditional direct standardization, which can adjust

for age and gender and some other discrete covariates (Carriere and Roos 1994, 1997). While we demonstrate a model-based method that can also deal with continuous covariates using large data files, this latter method is especially useful for the purpose of statistical hypothesis testing with correlated longitudinal data.

Innovative approaches are required to handle very large data sets efficiently or to analyze such nonbinary data as total days of hospitalization and discharges; recurrent events—events that may occur more than once in a specified time period to a given individual in the population of interest—can present problems. Portraying any kind of statistical distribution has proved difficult for such data, while the computer-intensive iterative procedures required for estimating parameters for nonlinear models are often not feasible to perform in large data files. Because only up to 20 percent of the population uses hospital resources in any given year, "no events" by over 80 percent of the population have to be dealt with.

Estimating and testing for parameters of interest can be done by dealing with those problems that frequently are encountered in analyzing longitudinal data: the correlation among repeated measurements, the time dependence of the event, the overdispersed distribution of the data, and the differing length of follow-up (Liang and Zeger 1986; Diggle, Liang, and Zeger 1995). In longitudinal analyses, failure to account for the correlation can result in smaller standard errors when between-unit comparisons (such as among income quintiles) are being made. But ignoring the correlation can result in larger standard errors when within-unit comparisons (such as among years) are being made.

We first distinguish our goals in this longitudinal study. Usually a longitudinal study observes subjects over time and traces person-specific change or growth. However, person-specific correlation and change is less relevant for studying change in health care utilization in the population over time; furthermore, the correlations over time within the Winnipeg population were quite small and assumed to be negligible. For example, the largest person-specific correlation over time based on 227,000 individuals admitted in the fiscal year 1989 was slightly positive at .01 in 1990 and became negative to −.12 in 1996; similar correlations for length of stay were .014 in 1990 and −.11 in 1996, indicating a weakly decreasing trend over time.

Given negligible patient-specific correlations, we seek to create subgroups whose units are highly homogeneous within each subgroup. For example, utilization by 0–2-year-olds in one year is likely to be similar to the utilization by 0–2-year-olds in the other years, and so on. This grouping

definition creates subgroups that are highly alike with respect to utilization within each group. Summarizing the data in this way does not result in much loss of information about utilization. We then concentrate on year-to-year utilization patterns of each subgroup. To accommodate for the correlation among longitudinal data on hospital utilization, we consider an exchangeable correlation model (an equal correlation in utilization patterns between any two years). No other form of correlation structure appears appropriate for the summary measures defined.

We adjust for all covariates as main effects as well as other significant interaction effects, because systematic fluctuations over time among the income quintiles may lead to interaction effects. Individual-level characteristics such as age and gender are among the most important predictors of health behavior (Roos, Black, Frohlich, et al. 1996).

In applying the GEE method to this very large data set, we stratify the individuals by age, gender, income quintile, and time of measurement, recording the average of events (average number of hospital discharges, average hospital days, and so forth). We use each consecutive age $(0, 1, 2, \ldots$ 84, and 85+) in the stratification, taking the last day of each fiscal year as the observation time for that year. An individual has to have been resident in Winnipeg for the entire year to be counted. Hence, the data form a rectangular structure with no missing values. Dividing the data into eight fiscal years generates eight measurements for each age/gender/income group with the number of individuals in each uncorrelated group varying from 50 to 2,159. Since the GEE method depends on an asymptotic theory, the number of strata (clusters) must be large enough both to ensure valid inferences and to retain all pertinent covariates. We analyze 172 subgroups in each of five income groups for a total of 860 clusters; each of these clusters is defined by age/gender/income and is assumed to be independent of other clusters. With no other individual-specific covariates being used for adjustment, averaging the data at each stratum causes no substantial loss of information.

Our hypotheses involve comparing utilization in 1989–1991 with that in 1992–1996; appropriate tests can be generated for the rates of events (individuals hospitalized, hospital discharges, days of hospitalization, and so forth; see appendix). The relative rates (RR) of decrease or increase in post-reform years compared to the pre-reform years are also reported. The GEE method enables testing of overall changes and, if the overall significant changes are indicated, changes for each income group also, based on five- and one degree of freedom (df) chi-square tests, respectively. The yearly trends (slopes) in ratios between poorest and wealthiest income quintile neighbor-

hoods were also examined using a one df chi-square test for a linear trend contrast (Snedecor and Cochran 1980). The Bonferroni adjustment is used to control the overall significance level when making multiple comparisons (Snedecor and Cochran 1980) over income groups.

RESULTS

We report our findings based on the analysis of the summary measures, acknowledging that the ability to draw detailed conclusions could depend on various micro-level factors that might not have been captured. However, our statistical strategy has relied on aggregating individual information without significantly altering the individual relationships in the data; our results and interpretations should be quite tenable.

Tables 2, 3, and 4 present hospital use characteristics of the Winnipeg population across the eight years and five income groups (detailed results are available from the senior author). Both relative ratios of the rate of the poorest income group (Q1) to that of the wealthiest income group (Q5) in each year are reported as well as three ratios for utilization: 1991/1989 (before health reform), 1996/1992 (during health reform), and 1996/1989 (over the eight years) (columns at the right of tables). The far right-hand column quantifies the relative changes in rates in post-reform years compared to pre-reform years.

Comparing the 1996/1989 ratios and the pre-post relative rates in Table 2 shows the reduction in days of hospitalization to be more dramatic than the fall either in the number of individuals hospitalized or in the number of hospital discharges.

Short-Stay Care Rates

Individuals Hospitalized for Short-Stay Care. Adjusting for both covariates and subgroup population sizes provides smoothed predictions of the rate of individuals hospitalized for short stays. Variations, both across income groups and across fiscal years, were found to be statistically significant, with all p-values less than .001. However, rates did not increase or decrease monotonically across the fiscal years and the income groups. As seen in Table 2, at least part of this variation is attributable to 1990, the year of the nurses' strike. Compared to the minor declines between 1989 and 1991, significant changes in rates of hospitalization in post-reform years 1992–1996 are noted in all income groups ($p < .001$); the relative rates of decrease ranged from .89 to .91 in post-reform years. Over the eight-year period, the rate of hospitalized

Table 2:   Hospital Use Characteristics of the Winnipeg Population (1989–1996)

Short-Stay Hospital Use†	Quintile	Fiscal Year								Ratios			Relative Rate
		1989	1990	1991	1992	1993	1994	1995	1996	91/89	96/92	96/89	Pre-Post
Individuals Hospitalized‡													
Relative affluence of	Q1*	92.62	86.76	89.44	88.10	85.14	82.82	80.07	78.42	0.97	0.89	0.85	0.91
neighborhood	Q2*	76.38	73.37	75.57	72.52	70.55	68.69	65.72	63.92	0.99	0.88	0.84	0.90
of residence	Q3*	70.42	69.64	70.76	67.96	66.68	63.43	62.30	60.35	1.00	0.89	0.86	0.90
	Q4*	67.38	63.70	66.03	62.48	61.08	58.65	57.59	54.90	0.98	0.88	0.81	0.89
	Q5*	60.96	56.24	59.95	57.13	54.58	52.98	51.60	49.16	0.98	0.86	0.81	0.89
	Overall	73.52	69.91	72.32	69.61	67.57	65.28	63.43	61.82	0.98	0.88	0.83	0.90
	Ratio: Q1/Q5§	1.52	1.54	1.49	1.54	1.56	1.56	1.55	1.60	0.98	1.03	1.05	1.03
Discharges‡													
Relative affluence of	Q1*	125.10	115.30	119.40	120.50	116.10	113.40	109.10	106.60	0.95	0.88	0.85	0.94
neighborhood	Q2*	102.00	97.32	101.30	97.63	95.82	93.50	89.28	87.70	0.99	0.90	0.86	0.93
of residence	Q3*	93.02	92.87	96.26	93.02	89.92	87.15	86.67	84.89	1.03	0.91	0.91	0.94
	Q4*	92.43	85.29	89.33	83.21	85.69	80.56	82.42	77.77	0.97	0.93	0.84	0.92
	Q5*	85.02	76.20	81.52	77.68	77.36	76.39	74.11	70.26	0.96	0.90	0.83	0.93
	Overall	99.48	93.35	97.51	94.37	92.94	90.16	88.27	85.42	0.98	0.91	0.86	0.93
	Ratio: Q1/Q5§	1.47	1.51	1.46	1.55	1.50	1.48	1.47	1.52	1.00	0.98	1.03	1.01
Days of Hospitalization‡													
Relative affluence of	Q1*	1101.00	1021.00	1003.00	977.70	885.10	878.30	833.10	811.30	0.91	0.83	0.74	0.84
neighborhood	Q2*	764.60	717.80	745.80	684.60	629.00	584.40	565.60	552.20	0.98	0.81	0.72	0.81
of residence	Q3*	666.80	628.30	626.30	596.90	551.50	506.20	485.10	454.20	0.94	0.76	0.68	0.81
	Q4*	597.60	575.40	570.60	495.20	441.60	402.80	385.20	387.10	0.95	0.78	0.65	0.72
	Q5*	499.80	454.70	484.70	438.20	366.30	359.70	344.30	346.30	0.97	0.79	0.69	0.77
	Overall	725.47	678.94	685.65	638.09	574.20	545.89	522.24	509.88	0.95	0.80	0.70	0.79
	Ratio: Q1/Q5**	2.20	2.25	2.07	2.23	2.42	2.44	2.42	2.34	0.94	1.05	1.06	1.09

*Pre- vs. post-health reform test significance level <.0001.
**Trend test significance level < .01.
†Only stays of 59 or fewer days are included.
‡Rate per 1,000 residents; adjusted using the parametric method.
§Q1/Q5: Q1 = poorest; Q5 = wealthiest.

persons has decreased significantly (1996/1989 ratios). However, the Q1/Q5 ratio that computes utilization by the poor relative to the affluent ranged from 1.49 to 1.60, suggesting stability over the years. The relative decrease in utilization observed in 1996/1989 ratios was similar for both Q1 and Q5.

Discharges from Short-Stay Care. As with the individuals hospitalized, variations in discharges from short-stay care across income groups and fiscal years were statistically significant ($p < .0001$). A generally decreasing pattern of hospital utilization over the eight fiscal years was observed, with decreases greater in the post-reform period ($p < .001$); the relative rates of decrease ranged from .92 to .94 in post-reform years. However, the Q1/Q5 ratios remained stable and appeared unaffected by the mandated reduction in bed supply.

Days of Hospitalization for Short-Stay Care. Short-stay hospital days per 1,000 population were examined using a logarithmic transformation to make the distribution of rates approximately normal. Short stays per 1,000 were significantly reduced in the reform years (1992–1996) compared to the previous three years ($p < .0001$), both overall and in each income group; the relative rates of decrease ranged from .72 to .84. However, the pattern of change over time significantly differed by income groups ($p < .0001$). Compared to the wealthiest income group, the poorest income group utilized significantly more hospital days in all years, confirming the cross-sectional findings of Roos and Mustard (1997). The Q1/Q5 ratios for individuals hospitalized and hospital discharges are considerably less than the Q1/Q5 ratio for the days of hospitalization (Table 2); these differences occur primarily because less affluent individuals are likely to experience longer lengths of stay. Utilization by the relatively poor declined significantly less than did use by the relatively affluent over the study years ($\chi^2 = 6.86$, $p < .01$).

Overall, rates of individuals hospitalized, hospital discharges, and length of hospital stay in all income quintiles declined more in the post-reform period than in the pre-reform period (the largest p-value $< .001$).

Utilization for Various Conditions and Procedures

To simplify presentation, Table 3 shows just quintiles 1, 3, and 5, and compares 1989 and 1996 for various conditions and procedures. Presenting the first and last years of data seemed appropriate, since the conditions and procedures showing overall declines (all except outpatient surgery and obstetric hospitalizations) dropped every year (with the exception of the 1990–1991 post-strike comparison) for almost all income groups. All Q1/Q5 ratios for these conditions and procedures remained stable over the years.

Table 3:   Hospitalization Rates by Income Quintile for Various Conditions and Procedures (1989 and 1996)

Variable	Quintile	Fiscal Year		Ratio: 96/89	Relative Rate Pre- Post
		1989	1996		
High-Variation	Q1*	36.26	23.24	0.64	0.77
Surgical Procedures	Q3*	31.40	21.07	0.67	0.80
	Q5*	28.30	19.16	0.68	0.81
	Overall	31.53	21.16	0.67	0.80
	Ratio: Q1/Q5§	1.28	1.21	0.95	0.94
Outpatient Surgery	Q1*	30.32	45.00	1.48	1.30
	Q3*	30.90	47.71	1.54	1.34
	Q5*	29.21	47.00	1.61	1.33
	Overall	30.70	46.77	1.52	1.31
	Ratio: Q1/Q5§**	1.04	0.96	0.92	0.98
High-Variation	Q1	56.33	50.84	0.90	0.96
Medical Conditions	Q3	37.56	37.87	1.01	0.96
	Q5	34.90	29.66	0.85	0.95
	Overall	42.24	38.14	0.90	0.95
	Ratio: Q1/Q5§	1.61	1.71	1.06	1.01
Low-Variation	Q1	6.58	5.77	0.88	1.03
Conditions	Q3	5.60	5.20	0.93	0.97
	Q5	4.96	4.87	0.98	1.00
	Overall	5.61	5.19	0.93	1.00
	Ratio: Q1/Q5§	1.33	1.18	0.89	1.03
Obstetrical	Q1***	101.90	101.50	1.00	1.04
Hospitalizations††	Q3	71.26	77.57	1.09	1.02
	Q5	65.12	59.54	0.91	0.94
	Overall	78.22	78.25	1.00	1.01
	Ratio: Q1/Q5§	1.56	1.70	1.09	1.10

Notes: Rate per 1,000 residents; adjusted using the parametric method. Only stays of 59 or fewer days are included.

*Pre- vs. post-health reform test significance level < .0001.

**Trend test significance level < .05.

***Pre- vs. post-health reform test significance level < .01.

†Rate per 1,000 residents; adjusted using the parametric method.

††Denominator included only women ages 15-44.

§Q1/Q5: Q1 = poorest; Q5 = wealthiest.

Differences in high-variation surgical procedures by income quintile were relatively small in either year (a Q1/Q5 ratio of 1.28 and 1.21, respectively, for 1989 and 1996). However, the number of such inpatient procedures dropped markedly, with the 1996 rates ranging between 64 and 68 percent of the 1989 rates. Although not shown in detail in Table 3, the significant decrease during post-reform years was experienced among all income groups (all p-values $< .0001$), RR ranging from .77 to .81. However, utilization by relatively poor individuals remained stable when compared to that by relatively affluent persons over the study years.

The dramatic increase in outpatient surgery between 1989 and 1996 proceeded from an initially small base, with rates for the Q1 quintile increasing by 48 percent and those for the Q5 quintile growing by 61 percent. Overall, this growth resulted in 52 percent more outpatient surgery being done in 1996 than in 1989. The increase during post-reform years was indicated with RR ranging from 1.30 to 1.34 (all p-values $< .0001$). However, the increase in outpatient surgery by the relatively poor was significantly less than that by the relatively affluent over the study years ($\chi^2 = 3.87$, $p < .05$).

Hospitalizations for high-variation medical conditions, low-variation conditions, and obstetrical care changed relatively little between 1989 and 1996. Although not evident from Table 3, the hospitalization for high-variation medical conditions was significant for Q2 and Q4 with RR $= .93$ and .92, respectively. The lowest income group did have significantly more births (RR $= 1.04$, $p < .01$) during the post-reform period (1992–1996) compared to the pre-reform period (1989–1991), while those for other income groups generally decreased or remained stable.

Conditions for Which Hospitalization Might Not Be Necessary

Three frameworks have been proposed to specify conditions for which "medical treatment is believed to be effective in preventing the condition, finding and treating the condition in an early phase to avoid major consequences, or treating the condition in a late phase, thereby avoiding death or disability" (Roos and Mustard 1997). These frameworks have been designated ambulatory care–sensitive conditions (Billings, Zeitel, Lukomnik, et al. 1993), avoidable hospitalizations (Weissman, Gatsonis, and Epstein 1992), and conditions amenable to medical treatment (Desmeules and Semenciw 1991).

Overall, the 1989–1996 period showed a moderate decline in hospital discharges for conditions for which such hospitalization might not be necessary (Table 4). Depending on the measures chosen, the 1996/1989 ratios varied between .79 and 1.06 with no consistent differences across

income quintiles; the rate for avoidable hospitalizations was an exception, decreasing significantly in the post-reform years (this, although not shown, was attributable to the significantly low rate for the Q2 in the post-reform period (RR = .89, p < .01).

The statistically significant declines were in general accord with those for high-variation medical conditions and for low-variation conditions. The health reform process has not resulted in special attention being given to these conditions, where hospitalization might have been avoided.

DISCUSSION

Methodological Advances

Our statistical strategy has relied on "appropriate aggregation," on aggregating individual information to permit the use of powerful statistics without significantly altering the individual relationships in the data. Such deliberate aggregation of the data is the counterpoint to efforts to estimate effects on

Table 4:   Rates by Income Quintile, Conditions for Which Hospitalizations Might Not Be Necessary

Variable	Quintile	Fiscal Year		Ratio: 96/89	Relative Rate Pre-Post
		1989	1996		
Ambulatory Care–	Q1	14.84	14.17	0.95	1.02
Sensitive Conditions	Q3	9.84	10.44	1.06	1.00
	Q5	8.24	7.83	0.95	1.01
	Overall	10.64	10.41	0.98	0.99
	Ratio: Q1/Q5[§]	1.80	1.81	1.00	1.01
Conditions Amenable	Q1	18.34	16.70	0.91	1.02
to Medical Treatment	Q3	13.30	12.20	0.92	0.98
	Q5	11.58	9.10	0.79	0.96
	Overall	13.93	12.18	0.87	0.98
	Ratio: Q1/Q5[§]	1.58	1.84	1.16	1.06
Avoidable	Q1	7.05	6.80	0.96	1.05
Hospitalizations	Q3	5.14	4.60	0.89	0.90
	Q5	3.87	3.89	1.01	1.02
	Overall	5.19	4.90	0.94	0.95
	Ratio: Q1/Q5[§]	1.82	1.75	0.96	1.03

Notes: Rate per 1,000 residents; adjusted using the parametric method. Only stays of 59 or fewer days are included.

[§]Q1/Q5: Q1 = poorest; Q5 = wealthiest.

individuals from aggregate or ecological data. Person-level data can be redefined for studying changes in utilization over time because the regression coefficients have meaning for the population rather than for any specific individual. Use of population-averaged responses in the statistical models describes ways in which average events across the population change according to their covariates when the focus is the population rather than any specific individual. We note, however, that our strategy would fail if the aggregation was "inappropriate" in ignoring important micro-level factors. Furthermore, while person-specific responses provide additional information about individuals, their modeling would involve additional assumptions about the correlation structure over time and is correct only when such assumptions are valid.

We have extended the cross-sectional method of evaluating small areas or group variations to consider rates of events observed over time. To account effectively for the dependent structure in the longitudinal data, subgroup responses were considered with an exchangeable correlation structure for repeated measurements in a quasi-likelihood framework; these models require that only the mean and variance functions be specified without making full distributional assumptions.

Our approach goes beyond simple age-gender adjustment in supporting analyses of appropriate age-gender specific data, dealing with changes in the overall size of the population served, and accommodating potential changes in its health status over time (see the appendix). The data need only be divided into subgroups fine enough to include such key characteristics as age, gender, and income level of residential area and to satisfy the large-sample assumption in the approach. The approach can be easily extended to include other important micro-level variables. Our strategy adjusts for both continuous and dichotomous covariates while accommodating longitudinal measurements for each subgroup.

The ability to test for significance over time and across space is important for assessing the effects of major system changes (such as a mandated reduction in bed supply). This article has concentrated on changes in utilization over time and socioeconomic groups (defined using neighborhoods, which are very small areas). However, the method of aggregation and significance testing is of wide general applicability for studying areas of different sizes over multiple points in time. In the Canadian context, the number of strata can be expanded to include more provinces. As such information from other provinces becomes available, comparisons of the effects of different strategies of reform across provinces will become more systematic. In the U.S.

context, our approach is particularly suitable for the large data sets involved in generating hospital report cards and, more generally, in assessing changes over time using state, Medicare, and Medicaid data (Hannan, Siu, Kumar, et al. 1995; Jollis and Romano 1998).

Health Reform

We have presented a picture of the Winnipeg hospital system as it was undergoing reform over the 1989–1996 period. Change has been gradual, although the beginning of "health reform" in 1992 certainly appears to have accelerated the process. The data show enduring differences in hospital utilization among Winnipeg residents of differing socioeconomic status as well as declines in fiscal 1992–1996. However, the significant interaction effects between income group and year indicate that the changes cannot be explained simply, and the nurses' strike in 1990 appears to have played a role in generating these interaction effects. The decrease in hospitalizations for low-variation conditions may indicate decreases in the incidence of these conditions, changes in practice patterns, changes in coding, or some combination of the three during the study period.

The relationship between surgery and hospitalization has been transformed with an overall increase in surgery that resulted from a decline in inpatient surgery coupled with the dramatic growth in outpatient surgery. Previous findings, that surgical care "appears remarkably unresponsive to the health needs of the population" (Roos and Mustard 1997), were based on the fact that rates for both high-variation surgical procedures and outpatient surgery varied relatively little with neighborhood income levels (our proxy for health status). Because the poor have been shown in a number of studies to need more care, Q1/Q5 ratios approximating unity for surgery have been seen as inequitable (Roos and Mustard 1997). From the perspective of health reform, the changes may raise a concern that the 1989–1996 increases in surgical rates are slightly less likely to affect the less affluent; we find that the Q1/Q5 ratio for outpatient surgery dropped from 1.04 (in 1989) to 0.96 (in 1996). This yearly decreasing trend in Q1/Q5 ratios was significant, indicating that the increase by relatively poor persons was significantly less than that by the relatively affluent population over the study years ($\chi^2 = 3.87$, $p < .05$), as expected.

In summary, change has been distributed across the different conditions, with a slight drop in medical hospitalizations and a slight increase in overall surgery, given the substitution of outpatient for inpatient procedures. Changes

in the Q1/Q5 ratio were small, even when they were statistically significant. The major changes—the decline in inpatient surgery, the growth in outpatient surgery, and the reduced number of hospital bed days—were necessitated by the 22 percent drop in acute care beds over the 1989–1996 period. These changes that accompanied Canadian health reform are important in that they do not closely correspond with what might have been suggested by the literature on variations in utilization (Wennberg and Gittelsohn 1982; Wennberg 1986; Fisher et al. 1994).

Drops in rates of surgical procedures have generally not occurred in Winnipeg, despite the fact that almost all procedures vary markedly across areas. Our detailed investigation reveals that the two Winnipeg teaching hospitals are disproportionately visited by residents of low-income neighborhoods. Given the large cuts in teaching hospital beds, our findings of little impact on the existing gradient of use by socioeconomic status are made even more striking.

Manitoba Health and the Manitoba College of Physicians and Surgeons have been unwilling to try to micro-manage utilization rates. Instead, cost efficiency has been increased by the shift to outpatient surgery. Surgical and medical hospitalizations are being handled very differently in the health care system. Given the frameworks that specify the medical conditions for which hospitalization might not be necessary, we tried to use these frameworks to predict the medical admissions which might have been reduced through a mandated reduction in bed supply. Our efforts did not prove fruitful; the magnitude of the reduction and its distribution among income quintiles varied considerably among frameworks.

APPENDIX

Generalized Estimating Equations (GEEs) Method

Let $\mathbf{y}_i = (y_{i1}, \cdots, y_{ir_i})'$ denote the vector of responses from each experimental unit $i = 1, \ldots, n$ measured at r_i different times (in our analysis, the $n = 860$ subgroups consist of the experimental units measured over $r_i = 8$ years for all i). Also let the $r_i \times p$ matrix $\mathbf{X}_i = (\mathbf{x}_{i1}, \cdots, \mathbf{x}_{ir_i})'$ with each $\mathbf{x}_{ij}$, a $p \times 1$ vector be the fixed p covariates associated with outcome y_{ij}, $j = 1, \cdots, r_i$. The y_{ij} are assumed to be a member of a generalized exponential family with a scale parameter ϕ and $\theta_{ij} = f(\mathbf{x}'_{ij}\boldsymbol{\beta})$, a function of the covariate vector $\mathbf{x}_{ij}$ and a vector of coefficients $\boldsymbol{\beta} = (\beta_1, \cdots, \beta_p)'$. The y_{ij} have mean

and variance denoted as μ_{ij} and ν_{ij}, respectively. Liang and Zeger (1986) considered generalized estimating equations (GEEs) to estimate regression parameter β by solving the equations:

$$\sum \mathbf{D}_i' \mathbf{V}_i^{-1} \mathbf{s}_i = 0$$

where $\mathbf{D}_i$ is the $r_i \times p$ matrix of derivatives $\partial \mu_i / \partial \beta$, $\mathbf{s}_i = \mathbf{y}_i - \mu_i$, with $\mu_i = (\mu_{i1}, \cdots, \mu_{ir_i})'$, and $\mathbf{V}_i$ is an $r_i \times r_i$ covariance matrix of $\mathbf{y}_i$. Then, by writing:

$$\text{cov}(\mathbf{y}_i) = \phi \mathbf{A}_i^{1/2} \rho_i(\boldsymbol{\alpha}) \mathbf{A}_i^{1/2}$$

where $\mathbf{A}_i = \text{diag}(\nu_{ij} / \phi)$ and $\rho_i(\boldsymbol{\alpha})$ is a working correlation matrix for each subgroup's multiple outcomes (i.e. longitudinal data), Liang and Zeger (1986) proposed a robust estimator of variance for the estimator $\hat{\beta}$, which was shown to be consistent even when ρ is misspecified, and demonstrated that solution of the GEEs produced consistent and asymptotically normally distributed estimates of the model parameters β. The variance of the solution $\hat{\beta}$ is estimated by:

$$\hat{\mathbf{C}}(\hat{\beta}) = \mathbf{H}_2^{-1} \mathbf{H}_1 \mathbf{H}_2^{-1}$$

where

$$\mathbf{H}_1 = \sum \mathbf{D}_i' \mathbf{V}_i^{-1} \mathbf{s}_i \mathbf{s}_i' \mathbf{V}_i^{-1} \mathbf{D}_i, \quad \text{and} \quad \mathbf{H}_2 = \sum \mathbf{D}_i' \mathbf{V}_i^{-1} \mathbf{D}_i.$$

Application of GEEs Method to Health Utilization Data

For each of the 860 clusters, we summarize their crude rate (y_{ij}) and the cluster size (n_{ij}) for the jth fiscal year from subgroup (cluster) i. This is done separately for each outcome. The choice of a link function will depend on the nature of the distribution of the outcome measure being analyzed. These are divided into three categories: (1) binary and dichotomous events (yes/no), such as individuals hospitalized; (2) counted events, which include all discharge data including discharges for surgical procedures (both inpatient and outpatient) and medical conditions; and (3) continuous data, such as days of hospitalization.

Note that this approach can accommodate the potential changes in health status over time as the covariate $\mathbf{x}_{ij}$ is allowed to vary across time of measurements and is specific to the subgroup from cluster i at fiscal year j. The predicted utilization measures from this approach were then summarized using the direct standardization method to control for the changes in the overall size and distribution of the population served.

1. Binary Outcomes. The logit link is most suitable for the binary event rate y_{ij} from the ith cluster, $i = 1, \ldots, n$, in the jth fiscal year, $j = 1, \ldots, J$, as:

$$\text{logit } E(y_{ij}) = \text{logit } (R_{ij}) = \mathbf{x}'_{ij}\boldsymbol{\beta} \quad \text{and} \quad \text{var}(y_{ij}) = R_{ij}(1 - R_{ij})\phi/n_{ij},$$

where $\mathbf{x}$ is the vector of predictor variables and n_{ij} is the size of the ijth subgroup. We used this approach for analyzing individuals hospitalized for short-stay care.

2. Counted Outcomes. Upon redefining the hospital discharge data as interval rate data, we may assume a Poisson model for the counted event rate y_{ij} with a logarithmic link:

$$\ln E(y_{ij}) = \ln R_{ij} = (\mathbf{x}'_{ij}\boldsymbol{\beta}) \quad \text{and} \quad \text{var}(y_{ij}) = R_{ij}\phi/n_{ij}.$$

The distribution of data on low incidence rates can be closely approximated by a Poisson model; concerns about violating the independence assumption of a Poisson distribution were shown to be minimal for events with low incidence rates (Carriere and Roos 1997). We used this approach for all outcomes reported here, except for individuals hospitalized and days of hospitalization.

3. Continuous Outcomes. For days of hospitalization, we may assume a normal model. Since these measures are likely to have highly skewed distributions, a suitable transformation may be necessary to generate more symmetrical distributions. We employ a logarithmic transformation on the event rates, y_{ij} with an identity link function:

$$E(\ln y_{ij}) = \ln R_{ij} = \mathbf{x}'_{ij}\boldsymbol{\beta} \quad \text{and} \quad \text{var}(\ln y_{ij}) = \phi\sigma^2_{ij}.$$

A weighted normal model can be considered by accommodating an auxiliary variable having any positive value, such as the population size or the inverse of an empirical variance of y_{ij}. In this analysis we used n_{ij} as the weight.

Hypothesis testing for contrasts of interest, for example, between combined rates for the five years 1992–1996 and combined rates for three previous years is also possible. We would need to specify a particular contrast of interest $\mathbf{R}$, which will be of $r \times q$ matrix for r sets of linear contrasts with $\boldsymbol{\beta}$ having q components. Then a robust test for a contrast is obtained by computing $(\mathbf{R}\hat{\boldsymbol{\beta}})'[\mathbf{R}\{\hat{C}(\hat{\boldsymbol{\beta}})\}\mathbf{R}']^{-1}\mathbf{R}\hat{\boldsymbol{\beta}}$, which is assumed to be distributed as χ^2 with r degrees of freedom. All analyses were done using SAS. For example, to test the equality of rates between pre- and post-health reform periods for income quintile group k, that is, for:

$$H_{0k} : 1/3(\mu_{89k} + \mu_{90k} + \mu_{91k}) = 1/5(\mu_{92k} + \mu_{93k} + \mu_{94k} + \mu_{95k} + \mu_{96k})$$

and if the fiscal year 1996 is the baseline, the $1 \times q$ vector $\mathbf{R}$ gets a value of $1/3$ for components relevant to income group k in years 89, 90, and 91, and gets a value of $-1/5$ for components relevant to income group k in years 92, 93, 94, 95, and zero otherwise. The test statistics is distributed as χ^2 with 1 degree of freedom. For simultaneous test for $k = 1, 2, 3, 4, 5$ income groups, the $\mathbf{R}$ collects all five vectors of similarly specified contrasts for each of the income groups and the test is carried out using χ^2 with 5 degrees of freedom. Usually, this simultaneous test is preceded by the individual income group test.

Similarly, for testing the linear contrast:

$$H_{0k} : -7\mu_{89k} - 5\mu_{90k} - 3\mu_{91k} - \mu_{92k} + \mu_{93k} + 3\mu_{94k} + 5\mu_{95k} + 7\mu_{96k} = 0$$

the $1 \times q$ vector $\mathbf{R}$ gets a value of $(-7, -5, -3, -1, 1, 3, 5)$ for components relevant to income group k in years 89, 90, 91, 92, 93, 94, and 95, respectively, and zero otherwise. The resulting test statistic has a χ^2 with 1 degree of freedom.

ACKNOWLEDGMENT

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